Metropolitan Earnings Inequality: Union and Government-Sector Employment Effects*

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Objective. This study examines the effects of union density and government-sector employment on earnings inequality in Metropolitan Statistical Areas (MSAs) in the United States. Methods. Data on 167 MSAs from the 2000 Census are analyzed using standard regression techniques. Four measures of Atkinson’s index (ε = 0.5, 1.0, 2.0, 3.0) are used as the measure of earnings inequality for full-time, year-round workers. Results. MSAs with greater union density and greater government-sector employment have lower earnings inequality. The progressive effect of union density is strongest for earners in the middle of the distribution and less beneficial for workers at the bottom of the distribution. Government employment is generally associated with lower levels of earnings inequality, but state and federal government employment have the strongest effects. Conclusion. Even in the late 1990s, unions and government-sector employment remain effective at reducing earnings inequality.

In a time of rising inequality, declining unionization, and a shift toward neoliberal approaches to government intervention in the economy, we ask if two institutional factors, unions and government-sector employment, remain effective at reducing earnings inequality across U.S. cities in the late 1990s. In the past, various studies have examined why economic inequality varies across metropolitan areas (Lorence, 1985; Nelson and Lorence, 1988; Brem, Durden, and Gaynor, 1989; Levernier, 1999; Hyclak, 2000; Madden, 2000; McCall, 2000). The focus of these studies has been concerned, mainly, with the impacts of human capital, demographic composition, and industrial structure. However, these analyses have largely neglected how institutional factors such as unions and government-sector employment impact earnings inequality. Thus, in the current study, we examine the impact of both union density and government employment on earnings

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SOCIAL SCIENCE QUARTERLY, Volume 86, Number 5, December 2005 ©2005 by the Southwestern Social Science Association

Unions and Earnings Inequality

Many observers have noted the decline in union membership in both the United States and abroad (Bowles, Gordon, and Weisskopf, 1990; Western, 1995; Farber and Western, 2001). Some analysts have found evidence that a substantial proportion of the increase in earnings inequality would not have occurred had union membership levels remained stable (Freeman, 1993; Fortin and Lemieux, 1997). Additionally, a long tradition has suggested that unions reduce wage inequality between white-collar and blue-collar workers (Freeman and Medoff, 1984; Lewis, 1986). Given the long-term decline in the proportion of the U.S. labor force in labor unions, one might speculate that unions have become less effective at reducing earnings inequality than they were in the past. In fact, Wallace, Leicht, and Raffalovich (1999) find that the effectiveness of unions at increasing labor’s share of income declined after the late 1970s. However, even if they have declined in effectiveness, there is still good theoretical reason to believe that unions have an impact on earnings inequality. Both neoclassical and institutional economists have theorized, but with different conclusions, about the potential effects of unions on earnings inequality.

Neoclassical theorists have reasoned that unions may increase earnings inequality (Friedman, 1962). Following neoclassical logic, a given wage is set by market forces under the conditions of perfect competition within general equilibrium. The demand for labor is an inverse function of the wage rate, while the supply of labor is a positive function of the wage rate. The intersection of these functions results in the equilibrium point of labor supplied and labor demanded, reflecting the equilibrium wage rate. Unions act to restrict the flow of labor into the unionized sector. Under this condition, the union can achieve a wage rate greater than the equilibrium wage. However, the flow of labor into nonunionized sectors will increase because the demand for labor in the unionized sector will decrease, given the higher union wage. The result of the increase in supply in the nonunionized sector is to push down wages in this sector, thereby increasing inequality because of the differences between the wages of unionized and nonunionized employees (Friedman, 1962). Additionally, workers who were already in advantaged positions in the earnings distribution are likely to benefit because employers have a greater demand for their labor.

Alternatively, researchers have used institutional approaches and argued that unions reduce inequality in pay (Freeman and Medoff, 1984; Hirsch and Addison, 1986; Dinardo and Lemieux, 1997). Proponents of this view have outlined several reasons why unions may reduce the inequality in the distribution of earnings.
Unions, as representatives of the local labor force, push for wage policies that benefit the majority of the workforce. In doing so, unions push for a single-rate-per-job policy. By implementing such a policy, unions compel employers to attach wages to jobs rather than attaching wages to individual workers based on individual supervisory decisions. Unions are premised on an ideology of egalitarianism and in order to maintain members’ confidence they must function efficiently to maximize the wages of all the union’s constituents. By standardizing rates of pay, unions act to greatly diminish the individual variation in wages among workers with similar levels of experience, seniority, and skill (Freeman and Medoff, 1984; Hirsch and Addison, 1986). The collective bargaining contracts pursued by unions reduce the number of pay categories and detract from the employer’s ability to differentiate among workers.

Marxist political economists have introduced a slightly more complex argument. In a capitalist society, the parameters of class conflict are largely shaped by state actors such that “unions cannot act in the interests of the working class as a whole, and in acting in their own interests they create divisions within the working class” (Rubin, 1988:555). In an early study of MSAs, Hyclak (1979) found that unions reduce inequality among working-class men and working-class African-American men, but not among working-class women. Studying the effect of union membership on family income inequality, Podgursky (1983) found the greatest progressive redistributive effect on families with union members and a negligible effect on all families (union and nonunion combined), suggesting that union effects are not universally redistributive. Alternatively, Freeman and Medoff (1984:159) found that the percentage of metropolitan statistical areas unionized positively correlated with higher wages among unorganized workers.

Sociologists and economists have also examined “threat effects” or the propensity for employers of nonunion workforces to increase wages in response to external local unionization threats (Leicht, 1989). In a study of seven Indianapolis manufacturing industries, unorganized workers’ wages in unionized plants increased with higher levels of unionization, while the wages of unorganized workers in nonunion plants decreased in response to external local unionization, but the effect was not significantly different from zero when plant size was controlled (Leicht, 1989; see also Freeman and Medoff, 1984:160). The evidence suggests that greater union density may reduce earnings inequality, but not for all workers across the earnings distribution.

Our review of the above theoretical and empirical literature leads us to test two hypotheses concerning the effect of union density on earnings inequality.

H1: MSAs with greater union density have lower levels of earnings inequality.

H2: The redistributive effect of union density is more beneficial to earners in the middle of the distribution than to earners closer to the bottom of the distribution.
Government Sector and Earnings Inequality

The distribution of earnings in the government sector is likely to be more compressed than earnings in the private sector for at least two reasons. First, government employers typically put a ceiling on the pay of workers at senior levels (Peters, 1985), while paying low-skilled workers more than their private-sector counterparts (Fogel and Lewin, 1974; Field and Keller, 1976; Durden and Schwarz-Miller, 1982; Peters, 1985). As an employer, the government can set the price of local labor and in so doing may be motivated by nonmarket considerations such as the goal of inequality reduction (Lobao and Hooks, 2003). In this sense, the government serves a “social equity function” (Lobao and Hooks, 2003:520). For example, some consider the higher wages paid to women working for the federal government compared to private-sector employers an “appropriate wage” because it essentially offsets private-sector gender discrimination in wages (Belman and Heywood, 1996:131). Government employers are also more likely to hire women and minorities than are private employers (Peters, 1985:252).

High rates of unionization and professional membership in the government sector also suggest that government employment is likely to compress the earnings distribution in a local labor market (Peters, 1985). The government is likely to utilize bureaucratic procedures to attach wages to different occupations. By putting a ceiling on wages of highly-skilled workers, paying low-skilled workers more than their private-sector counterparts, and having greater union density than the private sector, we should expect less earnings inequality in MSAs with higher levels of government-sector employment.

In addition to the theoretical considerations noted above, it is especially important to examine government effects on earnings inequality because government employment may be less effective at addressing inequality today than it was in the past (e.g., Zipp, 1994; Lobao and Hooks, 2003). That is, in the contemporary context of labor market deregulation, a test of the hypothesis that government employment reduces earnings inequality seems especially important. Given the studies reviewed above, we predict higher levels of government-sector employment are associated with lower levels of earnings inequality. These studies lead us to our third and final hypothesis.

H₃: MSAs with greater levels of government-sector employment have lower levels of earnings inequality.

Data and Methods

This study uses the Metropolitan Statistical Area (MSA) as the unit of analysis. This unit is appropriate as MSAs have long been used to approximate labor markets (e.g., Sakamoto, 1988; Levernier, 1999). In the
following analyses, we use data on 167 MSAs obtained from the 2000 Census of Population and Housing. With the exception of two variables, all data are extracted from Summary File 3 of the 2000 Census.

**Dependent Variables**

The dependent variable is earnings inequality, and it is estimated from a 20-category distribution of annual earnings for full-time, year-round workers (FTYR) age 16 and over found in Summary File 3 of the 2000 Census of Population and Housing (USCB, 2002). We examine FTYR workers because such workers have a strong attachment to the labor force, and by restricting the analysis to this subgroup we control for variation in hours and weeks worked.

For purposes of computation, we assume workers in the top category have earnings equal to twice the category’s lower bound. We estimated mean earnings in the three intervals below the top-coded category by solving for a Pareto multiplier and then multiplying the lower bound of each income category by this value (Parker and Fenwick, 1983). In the case of improbable Pareto multiplier estimates that produced income estimates outside the lower or upper boundary for a given category, we estimated that category’s income using linear interpolation. For earnings intervals below the top four, we use linear interpolation to estimate average income.

Our metric of inequality for our dependent variable is Atkinson’s measure. We chose this measure for two reasons: (1) it satisfies the principle of transfers (Allison, 1978) and (2) the Atkinson’s index can be adjusted to be more or less sensitive to the bottom of the income distribution. Other measures, such as the Gini index, are known to be overly sensitive to differences at the middle of the income distribution, and measures that are more sensitive to the bottom of the distribution (e.g., the variance of logged income) can violate the principle of transfers (Allison, 1978; Schwartz and Winship, 1980). In light of these considerations, we implement Atkinson’s index as our measure of earnings inequality because it does not violate the principle of transfers and can be adjusted to be more or less sensitive to the bottom of the income distribution.

Atkinson’s index is expressed as:

$$ A = 1 - \left[ \frac{1}{N} \sum_{i=1}^{N} f_i \left( \frac{y_i}{\mu} \right)^{1-e} \right]^{\frac{1}{1-e}} $$

1USCB refers to the U.S. Census Bureau. Full-time, year-round workers are defined as those who worked 35 or more hours per week and 50 or more weeks per year (USCB, 2002).

2In a regression analysis, a positively sloped coefficient indicates a variable contributing to greater inequality while a negatively sloped coefficient indicates a variable contributing to lesser inequality.
where $e \neq 1$ and $e > 0$. The value of $e$ is known as the “inequality aversion parameter” and can be adjusted to make $A$ more or less sensitive to the bottom of the income distribution. Namely, greater values of $e$ increase the sensitivity of $A$ to differences at the bottom of the income distribution. The range of values implemented for $e$ usually span from 0.5 to 4.0 (e.g., Schwartz and Winship, 1980). When $e = 1.0$, the formula appropriate for computation of $A$ is:

$$A = 1 - \exp \left[ \frac{1}{N} \sum_{i=1}^{N} f_i \log \left( \frac{y_i}{\mu} \right) \right]$$ (2)

where $y_i$ is the estimated income for an income category, $f_i$ is the frequency of earners in the $i$th income category, and $\mu$ is the average income. We use four values of $e$ for Atkinson’s index: 0.5, 1.0, 2.0, and 3.0. As per the principle of transfers, higher values of $A$ represent higher levels of inequality and lower values represent lower levels of inequality.

The utility of Atkinson’s index can be illustrated with examples. Consider two MSAs: Bismarck, ND and Birmingham, AL. The Atkinson index, when $e = 0.5$, is 0.138 for Birmingham and 0.123 for Bismarck, indicating slightly less earnings inequality in Bismarck. However, when $e = 3.0$, the Atkinson index is 0.636 for Birmingham and 0.681 for Bismarck, indicating a greater level of earnings inequality in Bismarck than Birmingham. Such examples underscore the argument that it is important to consider the fact that Lorenz curves can cross at different points of an income distribution (Bishop, Formby, and Thistle, 1992), yielding ambiguous inequality rankings.

**Main Independent Variables**

Our measure of unions, union density, represents the percentage of employed wage and salary workers age 16 and over who are members of a labor union (Hirsch and Macpherson, 2003). Hirsch and Macpherson derived this measure from various CPS data sets to obtain accurate and representative estimates of union membership across MSAs. Since the Census Bureau asks respondents for income in the previous year (i.e., 1999), we use union-density information for 1998 to account for any potential lagged effect.

We employ three measures of government employment: federal government employment, state government employment, and local government employment. Each represents the percentage of workers age 16 and over who are classified as federal, state, and local government workers, respectively (USCB, 2002). These three variables measure government employment at all possible levels, following Lobao and Hooks (2003). A log transformation
Control Variables

We control for human capital, industrial structure, and other measures of labor market conditions. We measure industrial specialization with a Herfindahl index of employment concentration computed on information from 20 different industries reported in the 2000 Census. Greater Herfindahl scores reflect greater industry-employment concentration within a city and should be associated with lower inequality (Lorence, 1985; Diamond and Simon, 1990). Mean establishment size is the average establishment size and was computed from information in the CENSTATS database. Greater average establishment size should be associated with less earnings inequality (Lorence, 1985). To hold constant labor market conditions, we control for the logged unemployment rate with the expectation that greater levels of unemployment are associated with lower mean earnings and likely greater earnings inequality (Blanchflower and Oswald, 1994).

The presence of a large immigrant population is likely to enhance earnings inequality because it has been found to increase wage inequality (McCall, 2000). Therefore, we control for the presence of immigrants with percent immigrant, the percentage of the population that is foreign born. To control for MSA racial composition, we computed the percentage of the population that is not white, percent nonwhite, expecting a positive link to earnings inequality.

Human capital is controlled with three variables. We computed an entropy index of educational inequality, dispersion in education, for the dividedness in educational attainment of the population age 25 and over (Jacobs, 1985; Nielsen and Alderson, 1997). We also include percent college educated, which is the percent age 25 and over with a B.A. or higher. To control for inequality in experience, we include dispersion in age, an entropy index of dividedness in age to proxy the inequality in experience among those age 25 through 64. Each of these three human capital measures should be positively related to earnings inequality. Lastly, we include controls for three broad Census regions to account for possible region fixed effects.

For purposes of estimation we estimate standard ordinary least squares (OLS) equations for each of the four measures of inequality as a dependent variable and the aforementioned independent measures as regressors. In

One reader noted that since union density has increased in the government sector in recent years, this measure may be collinear with government-sector employment. The zero-order correlation between union density and each level of government-sector employment is rather low: −0.187 (federal government), −0.086 (state government), and 0.098 (local government). Additionally, we attempted a control for logged population size but it was not statistically significant.

All estimates were obtained with R 2.0.1 statistical software (Faraway, 2005). Since the data set is relatively small, we examined various regression diagnostics, including leverage
Table 2, we present the OLS estimates and the “HC3” standard errors fitting the same equation to Atkinson’s index at each level of inequality aversion.

Results

Table 1 presents the descriptive statistics and variable descriptions for all measures used in the analysis. Table 2 presents the results of our empirical estimates. Each model is fit to Atkinson’s index at a different level of inequality aversion, indicated by the parameter e. Higher values of e signify a metric increasingly sensitive to the lower tail of the distribution. Taken as a whole, the $R^2$’s range from 0.584 to 0.782, which is consistent with previous research.

Examining the control variables, industrial specialization is associated with lower levels of earnings inequality, in line with recent research (Wheel-er, 2004). Mean establishment size is associated with lower earnings inequality for workers closer to the bottom of the distribution. Surprisingly, the percent immigrant is not associated with earnings inequality, and for earners plots and jackknife residuals, and did not find any unusual or influential observations. Additionally, in supplementary analyses we estimated the models with an iteratively re-weighted least squares algorithm. These robust regression estimates were nearly equal to OLS estimates. Therefore, we use standard OLS regression to estimate parameters, along with “HC3” standard errors to account for potential heteroskedasticity (Long and Ervin, 2000). In addition, we used variance-inflation factors and eigenvalues to check for multicollinearity and these measures indicated that it is not a problem in this analysis.
at the bottom of the distribution it is actually associated with lower levels of earnings inequality. Unemployment is positively associated with earnings inequality. Consistent with the human capital model, greater dispersion in education and age are both associated with greater levels of earnings inequality across MSAs. The percent college educated is positively associated with earnings inequality for earners across the distribution. The percent of the population nonwhite is not statistically significant in any of the models.

Hypothesis 1 predicts that greater MSA union density is associated with lower levels of earnings inequality. Consistent with this hypothesis, the coefficients for union density are negatively sloped for both Models 1 and 2 in Table 2.

### TABLE 2

Union and Government-Sector Effects on Earnings Inequality (Atkinson \times 100) Among Full-Time Year Round Workers in 167 MSAs, 1999

<table>
<thead>
<tr>
<th></th>
<th>1</th>
<th>2</th>
<th>3</th>
<th>4</th>
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</thead>
<tbody>
<tr>
<td></td>
<td>e = 0.5</td>
<td>e = 1.0</td>
<td>e = 2.0</td>
<td>e = 3.0</td>
</tr>
<tr>
<td>Union density</td>
<td>-0.027*</td>
<td>-0.030</td>
<td>0.019</td>
<td>0.105*</td>
</tr>
<tr>
<td></td>
<td>(0.013)</td>
<td>(0.019)</td>
<td>(0.033)</td>
<td>(0.045)</td>
</tr>
<tr>
<td>Federal government</td>
<td>-0.444***</td>
<td>-0.687***</td>
<td>-1.089**</td>
<td>-1.245**</td>
</tr>
<tr>
<td>employment log</td>
<td>(0.101)</td>
<td>(0.173)</td>
<td>(0.368)</td>
<td>(0.476)</td>
</tr>
<tr>
<td>State government</td>
<td>-0.657***</td>
<td>-1.097***</td>
<td>-1.784***</td>
<td>-1.744**</td>
</tr>
<tr>
<td>employment log</td>
<td>(0.110)</td>
<td>(0.179)</td>
<td>(0.366)</td>
<td>(0.511)</td>
</tr>
<tr>
<td>Local government</td>
<td>-0.126*</td>
<td>-0.168*</td>
<td>-0.264</td>
<td>-0.272</td>
</tr>
<tr>
<td>employment</td>
<td>(0.051)</td>
<td>(0.082)</td>
<td>(0.142)</td>
<td>(0.183)</td>
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<tr>
<td>Control Variables</td>
<td></td>
<td></td>
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<td></td>
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<tr>
<td>Industrial</td>
<td>-0.002***</td>
<td>-0.003***</td>
<td>-0.004**</td>
<td>-0.004</td>
</tr>
<tr>
<td>specialization</td>
<td>(0.0003)</td>
<td>(0.0006)</td>
<td>(0.002)</td>
<td>(0.002)</td>
</tr>
<tr>
<td>Mean establishment</td>
<td>-0.060</td>
<td>-0.114</td>
<td>-0.428**</td>
<td>-0.607***</td>
</tr>
<tr>
<td>size</td>
<td>(0.038)</td>
<td>(0.062)</td>
<td>(0.126)</td>
<td>(0.163)</td>
</tr>
<tr>
<td>Percent immigrant</td>
<td>0.024</td>
<td>0.014</td>
<td>-0.067</td>
<td>-0.159**</td>
</tr>
<tr>
<td></td>
<td>(0.014)</td>
<td>(0.023)</td>
<td>(0.044)</td>
<td>(0.056)</td>
</tr>
<tr>
<td>Unemployment rate</td>
<td>1.046**</td>
<td>1.840**</td>
<td>3.295**</td>
<td>2.135</td>
</tr>
<tr>
<td>log</td>
<td>(0.332)</td>
<td>(0.569)</td>
<td>(1.194)</td>
<td>(1.778)</td>
</tr>
<tr>
<td>Dispersion in</td>
<td>4.885***</td>
<td>8.397***</td>
<td>16.187***</td>
<td>18.758***</td>
</tr>
<tr>
<td>education</td>
<td>(0.939)</td>
<td>(1.566)</td>
<td>(3.434)</td>
<td>(4.528)</td>
</tr>
<tr>
<td>Percent college</td>
<td>0.101***</td>
<td>0.146***</td>
<td>0.280***</td>
<td>0.280***</td>
</tr>
<tr>
<td>educated</td>
<td>(0.014)</td>
<td>(0.025)</td>
<td>(0.056)</td>
<td>(0.083)</td>
</tr>
<tr>
<td>Dispersion in age</td>
<td>7.268*</td>
<td>11.783*</td>
<td>8.654</td>
<td>0.947</td>
</tr>
<tr>
<td></td>
<td>(3.201)</td>
<td>(5.865)</td>
<td>(9.920)</td>
<td>(12.312)</td>
</tr>
<tr>
<td>Percent nonwhite</td>
<td>-0.011</td>
<td>-0.020</td>
<td>-0.026</td>
<td>-0.040</td>
</tr>
<tr>
<td></td>
<td>(0.006)</td>
<td>(0.010)</td>
<td>(0.018)</td>
<td>(0.022)</td>
</tr>
<tr>
<td>Constant</td>
<td>-11.768</td>
<td>-17.947</td>
<td>-7.854</td>
<td>28.599</td>
</tr>
<tr>
<td>$R^2$</td>
<td>0.782</td>
<td>0.745</td>
<td>0.684</td>
<td>0.584</td>
</tr>
</tbody>
</table>

*p < 0.05; **p < 0.01; ***p < 0.001 (two-tailed tests).

**Note:** Unstandardized (OLS) coefficients with heteroskedasticity-corrected (HC3) standard errors in parentheses. All models include three dummies for broad Census region. $N = 167$. 

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On average, each one percentage point increase in union density decreases Atkinson’s index by 0.027 ($e = 0.5$). This indicates that in MSAs with greater union density, the distribution of earnings between the highest-paid earners and earners in the middle of the distribution is more compressed.

Hypothesis 2 predicts that union density is less beneficial to earners at the bottom of the distribution than to workers in the middle of the distribution. Models 2, 3, and 4 are also fit to A, but with greater sensitivity to differences at the bottom of the distribution. In Models 2 ($e = 1.0$) and 3 ($e = 2.0$), the union-density slope estimates are not statistically different than zero, indicating that the union effect diminishes as more weight is placed on the bottom of the distribution.

Consistent with the idea that unions are more beneficial to workers in the middle of the earnings distribution, we find a positive and statistically significant effect of union density on earnings inequality when the greatest weight is placed on the bottom of the distribution ($e = 3.0$). Thus, when we enhance the sensitivity of the inequality measure to the bottom of the distribution, we find that in MSAs with greater union density, earnings inequality is somewhat higher between workers at the lowest points on the income distribution and the rest of the distribution.

Hypothesis 3 predicts that MSAs with greater levels of government-sector employment have lower levels of earnings inequality. Consistent with this hypothesis, earnings inequality is inversely associated with government-sector employment at the federal, state, and local levels. Local government employment is associated with lower levels of earnings inequality for earners in the middle of the distribution, but like the union effect, its impact is less beneficial for earners closer to the bottom of the distribution (Model 4).

One way to interpret the size and importance of the effects is to simulate the change in the mean level of earnings inequality after a specified change in a given variable. The average expected values of Atkinson’s indices are: 12.66, 22.42, 41.18, and 63.69. Setting the union effect to zero, the expected mean values of Atkinson’s index change by $+2.45$, $+1.59$, $-0.55$, and $-1.93$ percent (for $e = 0.5$, 1.0, 2.0, and 3.0, respectively). Thus, in the absence of unions, mean MSA earnings inequality would be between 1.6 and 2.5 percent higher. When emphasis is placed on the very bottom of the distribution, the absence of unions would reduce inequality by approximately 2.0 percent (using the statistically significant estimate from Model 4 in Table 2).

We may perform the same simulations for the effects of government-sector employment at the mean at each level of inequality aversion. Setting the log of federal government employment to zero and holding all else constant at the sample mean, earnings inequality would be $+3.00$, $+2.62$, $+2.26$, and $+1.67$ percent higher at each level of inequality aversion ($e = 0.5$, 1.0, 2.0, and 3.0, respectively) in the absence of federal government employment. In the absence of state government employment, earnings inequality at the mean would be $+8.31$, $+7.83$, $+6.93$, and $+4.38$
(e = 0.5, 1.0, 2.0, and 3.0, respectively) percent higher. The absence of local government employment would increase earnings inequality at the sample mean by +6.78 and +5.12 percent (e = 0.5, 1.0, respectively).

The results are consistent with the patterns predicted by Hypotheses 1 through 3. Earnings inequality is inversely related to union density, the beneficial effect of union density is felt most strongly by the earners in the middle of the distribution, and government-sector employment has a relatively strong, progressive effect on earnings inequality.

Discussion and Conclusion

In a time of rising inequality, declining union membership, and increasingly neoliberal approaches to government, we find that unions and government-sector employment remain inversely correlated with earnings inequality across MSAs in the late 1990s. Previous studies of metropolitan earnings inequality have investigated demographic, industrial, and aggregate organizational impacts on earnings inequality. In the present study, we have extended this literature by examining the effects of union density and government-sector employment at the federal, state, and local levels.

We find that unions are more beneficial to workers in the middle of the earnings distribution than to workers closer to the bottom of the distribution. That is, in MSAs with higher union membership, the inequality in the distribution of earnings between earners at the top of the distribution and earners at the middle of the distribution is lower. As noted above, unions have been found to reduce inequality between white-collar and blue-collar workers (Freeman and Medoff, 1984). A plausible interpretation of our results is that MSAs with greater union density have lower inequality between highly-paid white-collar earners and unionized blue-collar workers. The union-induced improvement of blue-collar workers’ earnings brings them closer to white-collar workers and further away from unorganized workers and this is responsible for the positive link between union density and earnings inequality for earners at the bottom of the distribution. In accordance with the Marxist hypothesis (see Rubin, 1988), unions cannot act in the interests of the working class as a whole and this is indicated by the fact that their beneficial effects are not felt throughout the entire income distribution.

MSAs with higher levels of government-sector employment continue to have lower levels of earnings inequality. The strongest effects stem from state government employment and federal government employment. Local government employment has an impact, but it is not felt across the entire range of the MSA earnings distribution. In the absence of federal and state government employment, MSA earnings inequality would be substantially higher. As noted above, government employers typically put a ceiling on the wages of workers in their highest-paid categories, are more likely to attach
wages to jobs through bureaucratic pay scales, and the public sector has witnessed growth in union density in recent years. MSAs with a greater proportion employed in the government sector likely have lower inequality because the government sector has a more compressed distribution of earnings. The presence of a large government sector likely also has spillover effects into the larger labor market motivating other employers to pay more equitable wages. Although the trends of declining union membership and government-sector downsizing have been in place for at least three decades, unions and government employment remain effective at reducing earnings inequality across metropolitan areas.

REFERENCES


